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Vegetable oil market and biofuel policy: An asymmetric cointegration approach

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ABSTRACT

The present paper analyses the long-run relationship between vegetable oil prices and conventional diesel prices in the EU for the period 2005–2007. We applied recent developments in the threshold cointegration approach to investigate the presence of asymmetric dynamic adjusting processes between the prices of rapeseed oil, sunflower oil, and soybean oil, and the price of a mineral oil: diesel. The results presented suggest a two-regime threshold cointegration model only for the rapeseed oil–diesel price pair. Thus, the rapeseed oil price adjusts rapidly to its long-run equilibrium, determined by fossil diesel prices, but this adjustment is asymmetric: it differs if the divergence between the two prices is above or below a critical threshold. Consequently, rapeseed oil appears particularly exposed to external shocks deriving from global political scenarios, suggesting the reassessment of the high quota (80%) of EU biodiesel represented by this vegetable oil.

1. Introduction

In recent years there has been growing concern about oil price rises, fuel security and environmental issues, and this has led policy makers to adjust their energy portfolios, with a more diversified range of energy sources used, and paying increased attention to the biofuel sector. This choice in particular has a strategic internal relevance as biofuel represents a source of energy that can be produced internally in an energy-resources scarce geographical entity. The high level of biofuel production in the European Union (EU) (4.8 million tons of biodiesel produced during 2006, equivalent to 77% of world-wide biodiesel production — EBB, 2007) is very closely related to the policy framework that has been specifically designed within the EU, both in the supply and demand side.

Among the different policies implemented in the EU to stimulate biofuel production, we recall the directive 2003/96/EC on energy taxation, the energy crop premium and the non-food set-aside payment. In order to make biofuel prices competitive with respect to conventional fuels, the energy taxation directive gives EU Member States a legal framework to apply excise duty to renewable fuel. The energy crops premium, introduced by the Common Agricultural Policy reform of 2003, grants a carbon credit payment (45 €/ha) to crop growers. Farmers can also receive an additional decoupled payment, in line with the Blair House agreement, by cultivating non-food crops on set-aside land.

On the demand side, we recall the directive 2003/30/EC on the promotion of biofuel use, which fixed a share of biofuel blends equal to 2% of the overall consumption of gasoline and diesel in transport for the end of 2005, rising to 5.75% in 2010. Moreover, the European Commission recently proposed a 10% biofuel target by 2020 (COM (2008)19 final).

This political framework has had a great impact on production: from 2005, the first biofuel target year of the directive 2003/30/EC, to 2007, the EU-25 biodiesel production rose to 57%. In the same period biodiesel production capacity increased from 4.2 to 10.2 million tons, the equivalent of an increment of 142% (EBB, 2007).

One of the most important effects of the growing production of biofuel was the change in the nature of the link between agricultural commodity prices and energy market. Traditionally, energy prices have affected agricultural prices as part of production cost (agricultural cost, transportation, and marketing), while in recent years energy prices have influenced agricultural prices through the increasing use of agricultural commodities as raw material for biofuel production. As a consequence, an increase in energy prices stimulates the price-competitiveness of biofuel sources, leading to an increase in the demand of specific agricultural crops (e.g. feedstock), useful in biofuel production.

The recent strong increase in demand for vegetable oils for biodiesel production the EU recorded in recent years has therefore activated new relationships among food and non-food commodity prices, especially between vegetable oils and substitute goods such as crude oil and its derivatives. While the economic aspects of the biofuel sector have captured scholarly attention (for a comprehensive review,

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see Rajagopal and Zilberman, 2007), this changing price relation has received little attention in academic literature despite the relevant effects on both the energy and the agricultural sector, and the present article aims to address this issue.

Within the economic literature, the approaches used to study price relationships follow different conceptual frameworks. One of the most widely used is based on the analysis of price transmission mechanisms within the context of market integration (e.g. Ravallion, 1986; Sexton et al., 1991; Zanias, 1993; Blauch, 1997) and the law of one price (e.g. Ardeni, 1989; Baffes, 1991). Within this framework, some *structural* approaches support the necessity of macroeconomic factors in understanding price movements of a commodity. On the other hand, *non-structural* studies consider those factors determining transmission as external prior information derived from theoretical propositions assumed in the variable selection rather than the outcome of a theoretical framework. In this work, we follow a non-structural approach, frequently used in price transmission analysis, along with a novel time series mechanism, to investigate price relations between vegetable oil and diesel.

In the existing literature, several studies analysed vegetable oil price relationships (Duncker, 1977; Labys, 1977; Griffith and Meilke, 1979; In and Inder, 1997; Owen et al., 1997), but just recently Yu et al. (2006) and Campiche et al. (2007) have studied the interdependence of vegetable and mineral oil prices. Yu et al. (2006), using time series mechanism relative to the period 1999–2006, found a long-run relationship between different vegetable oils and crude oil prices. Furthermore, although crude oil shocks had no significant effect on vegetable oil prices, the paper highlights that the influence of crude oil prices on edible oils will grow if high oil prices continue and edible oils become an increasing source of biodiesel. In line with the main results of the abovementioned study are Campiche et al.'s (2007) findings: splitting the time sample into two periods, 2003–2005 and 2006–2007, they draw attention to the increasing tendency for price change in petroleum prices to correspond to price changes in selected agricultural commodities.

Within the debate on market tradeoff between crude oil prices and edible oils, the aim of this paper is twofold. Firstly, it analyses the degree and the nature of price linkages between vegetable oil prices (rapeseed, sunflower seed and soybean oil) and fossil diesel prices during the period 2005 to 2007 after the implementation of several biofuel policies to support biodiesel production in the EU. Secondly, the paper investigates the possible asymmetric nature of price relationships between diesel and vegetable oils. Asymmetric price transmission assumes that prices adjustments are not homogeneous with respect to internal or external characteristics of the system and represents one important issue in the economic literature (Frey and Manera, 2007; Meyer and von Cramon-Taubadel, 2004). In the specificity of our work, recent biofuel policy could have changed the price response behaviour, making price movements in one market asymmetrically transmitted to the other market, where the asymmetry is manifested through different speeds of transmission in different regimes.

To explore these issues in our empirical analysis, we first test for the presence of a bivariate cointegration between the variables we include in our model. Subsequently, for the variables found to be cointegrated, the paper follows the techniques presented in the existing literature for investigating the adjustment processes, using a threshold vector error correction model (TVECM). While linear cointegration techniques have already been employed to model price linkages between vegetable oils and fuel, threshold cointegration has never been used to understand these links.[1] Studying asymmetric price transmission could give more information on the vulnerability of such markets, in particular in the presence of fuel shocks.

In the empirical analysis, we have chosen to investigate asymmetric market dynamics between several vegetable oils (which will be presented in the next sections) and fossil diesel only when prices show a common trend in the long run, i.e. are cointegrated. The rationale for our choice arises from our interest in the relationship between biofuel policy implementation and food/non-food commodity market, where the absence of long-run relations between prices implies the absence of tradeoffs between the markets in analysis, delivering limited insights for biofuel policy makers. On the other hand, if variables are found to be cointegrated, it will come as a consequence of the energy portfolio diversification obtained through the implementation of biofuel policy. As a result, there exists scope for investigating the adjustment process through a TVECM.

A better understanding of the oil price relation is of primary importance for both energy and agricultural policies. Within the EU energy policy framework, the great emphasis accorded to biomass-based energy binds policy maker to consider the effect of such policies on food prices and on the repercussion on environmental issues. This point is supported by our results, which find that soybean oil and sunflower seed oil show no long-run correlation with diesel prices. Furthermore, we find evidence of asymmetric movement between fossil diesel prices and rapeseed oil prices, with the first driving the second to its long-run equilibrium.

The remainder of the paper is composed of three parts, organized as follows. Section 2 reviews the threshold cointegration approach and recent developments in the related testing procedure. Section 3 outlines and presents the data sources and the results of the econometric analysis. The final section discusses our findings and concludes.

2. Theoretical issues

An extensive literature has applied cointegration techniques to investigate if long-run *equilibria* exist among prices. These traditional models assume that the adjustment process to maintain the *equilibrium* occurs in every time period. However, many situations, and, in particular the commodity price stabilization, are often characterized by discrete interventions. In other words, it is conceivable that it is only when the market price moves too far from the target price that economic agents intervene to bring the system back towards the *equilibrium*. Recently, two main classes of models have been proposed in the literature to characterize this kind of non-linear adjustment process. One class considers Markov-Switching Vector Error Correction models, assigning probabilities to the occurrence of different regimes (Hamilton, 1989; Krolzig, 1997). The second class is based on the ideas of Tong and Lim (1980) about Self-Exciting Autoregressive Model where the regimes that have occurred in the past and the present are known with certainty, though via statistical techniques. In this context, Balke and Fomby (1997) introduced the concept of "threshold cointegration", a feasible approach to allow the adjustment process to move differently in separate regimes. They hypothesized that this movement towards a long-run equilibrium may not occur in every time period, but only when the deviation from equilibrium exceeded a critical threshold, hence its name. In this sense, cointegration can be considered as a *global* characteristic of the series, while threshold behaviour as *local* and discrete characteristic.

Balke and Fomby's (1997) approach provides a framework for analysing the relationship between linear and non-linear (threshold) cointegration. In fact, the researcher faces four possible hypotheses: no cointegration and linear behaviour; no cointegration and non-linearity; cointegration and linearity; cointegration and non-linearity. Their proposed approach consists in breaking up the econometric analysis into two steps: firstly, they test for linear no cointegration against linear cointegration; subsequently, they test for linear cointegration against threshold cointegration. This approach is supported by the fact that many existing instruments for testing for cointegration can be of use even in the case of threshold cointegration (Hansen and Seo, 2002). The

[1] While this approach has not been used in testing the vegetable oil–mineral fuel relation, some papers have analysed the non-linear relation between ethanol and mineral oil. In particular, Rapsomanikis and Hallam (2006) and Balcombe and Rapsomanikis (2008) have applied a TVECM approach in the analysis of the sugar–ethanol–oil nexus in Brazil, while Serra et al. (2008) considered the corn–ethanol–oil case in the US.

Balke and Fomby's (1997) two-step procedure excludes the threshold-no cointegration hypothesis, as is frequently assumed in the empirical literature (e.g. Lo and Zivot, 2001; Goodwin and Piggott, 2001; Chung et al., 2005; Rapsomanikis and Hallam, 2006), which is not straightforward to test, requiring completely a different distribution theory and no single test can simultaneously verify the two conditions (Seo, 2006). Consequently, in the present exercise we follow the approach most commonly found in the literature, assuming that rejecting the linear no cointegration null hypothesis implies either linear or threshold cointegration. While acknowledging that this approach is suboptimal, we think it is effective and appropriate in answering our research question.

Analysing our work in more detail, the model we propose is a threshold vector error correction model (TVECM), with a threshold effect based on an error correction term.[2] In the case of two regimes, Balke and Fomby (1997) present a TVECM of order $\ell+1$ that takes the form:

$$\Delta x_t = \begin{Bmatrix} A_1' X_{t-1}(\beta) + u_t, & \text{if } w_{t-1}(\beta) \leq \gamma & \text{regime 1} \\ A_2' X_{t-1}(\beta) + u_t, & \text{if } w_{t-1}(\beta) > \gamma & \text{regime 2} \end{Bmatrix} \tag{1}$$

with:

$$X_{t-1}(\beta) = \begin{Bmatrix} 1 \\ w_{t-1}(\beta) \\ \Delta x_{t-1} \\ \Delta x_{t-2} \\ \vdots \\ \Delta x_{t-l} \end{Bmatrix} \tag{2}$$

where x_t is a p-dimensional time series I(1) cointegrated with one $(p\times 1)$ cointegrating vector β, $w_t(\beta)$ is the error correction term (ECT), u_t is the error term assumed to be an *iid* Gaussian sequence with a finite covariance matrix, A_1 and A_2 are matrices of coefficients describing the dynamics in each regime and finally γ is the threshold parameter. The values of w_{t-1} below or above the threshold γ allow the coefficients to switch between regimes 1 and 2; in particular, the estimated coefficients of w_{t-1} of each regime denote the different adjustment speeds of the series towards equilibrium.

Hansen and Seo (2002) indicated a method to estimate TVECM by maximum likelihood, which involves a joint grid search over the threshold parameter and cointegrating vector. In order to test for threshold cointegration, Tsay (1989, 1998) proposed non-parametric non-linearity tests, while Andrews (1993), Hansen (1996), Balke and Fomby (1997) and Lo and Zivot (2001) presented different methods based on Lagrange Multiplier (LM) statistics. More recently, Hansen and Seo (2002) developed two SupLM tests for a given or estimated β using a parametric bootstrap method to calculate asymptotic critical values with the respective p-values. The first test is denoted as:

$$\sup LM^0 = \sup_{\gamma_L \leq \gamma \leq \gamma_U} LM(\beta_0, \gamma)$$

and would be used when the true cointegrating vector β is known *a priori*. The second test is used when the true cointegrating vector $\tilde{\beta}$ is unknown and they denote this test statistic as:

$$\sup LM = \sup_{\gamma_L \leq \gamma \leq \gamma_U} LM(\tilde{\beta}, \gamma)$$

where $\tilde{\beta}$ is the null estimate of the cointegrating vector.

In these tests, the search region $[\gamma_L, \gamma_U]$ is set so that γ_L, is the π_0 percentile of $\tilde{w}_{t-1}$ [where: $\tilde{w}_{t-1} = w_{t-1}(\tilde{\beta})$], and γ_U is the $(1-\pi_0)$ percentile.[3]

3. Empirical analysis and results

For our analysis, we used weekly price data from January 2005 to November 2007 for rapeseed oil (RapOil), soybean oil (SoyOil), sunflower seed oil (SunOil) and fossil diesel (GasOil) in the EU. For the GasOil variable, we used the Rotterdam diesel prices taken from the Energy Information Administration (EIA), while the vegetable oil data were obtained from Oil World (ISTA Mielke GmbH); all prices are spot, expressed in US dollars/MT and converted to natural logarithms. In line with many empirical studies (e.g. Hu and Lin, 2008), we adjust the data to account for seasonal effects, transforming the data with monthly adjustments. In Table 1, seasonality was investigated with a standard *F* test, leading to a rejection of the null of no seasonal patterns for all the variables.

Considering cointegration as a *global* characteristic of the series while threshold behaviour as *local* characteristic (Balke and Fomby, 1997), we conduct the analysis as follows: first the degree of integration of the variables was tested by the Augmented Dickey–Fuller test (ADF), Philips–Perron test (PP) and Kwiatkowski–Phillips–Schmidt–Shin test (KPSS). Then, we test for cointegration and weak exogeneity between variables. Subsequently, we estimate cointegrating vectors and test for the presence of threshold cointegration. Finally, we estimated TVECM by the Hansen and Seo (2002) procedure.

Table 2 shows the results of the ADF and PP tests, where the *D* in front of every variable name indicates the differentiated series. It emerges that all the series are I(1) with and without trend. The KPSS test shows quite different results for GasOil and RapOil with trend and for SunOil without trend, but we can conclude that these series are I(1).

We investigated three relationships among the considered vegetable oils and GasOil, and tested for the presence of cointegration and the optimum lag length following two different approaches. The first refers to the well known Johansen procedure (Johansen, 1988, 1992; Johansen and Juselius, 1990), which involves a system based on a likelihood ratio test that contemplates two steps. In the first, the lag number of VAR representation is determined using information matrices based on Akaike (1973), Hannan and Quinn (1979) and Schwarz (1978) information criteria (IC). In the second step, given the optimal lag length, the cointegration rank is obtained through the trace test and the Maximum-Eigenvalue test. All cointegrating equations included an intercept, and regressions were estimated both with and without a linear trend in level data.

However, this two-step procedure might not be free from model specification problems in terms of tradeoffs between model parsimony and fit, given the fact that the true model is not frequently known (Wang and Bessler, 2005). Accordingly, to improve the performances of this procedure, a model selection method based on information criteria (one-step procedure) has been proposed as a valid alternative (Aznar and Salvador, 2002; Baltagi and Wang, 2007; Phillips and McFarland, 1997). Thus, the second approach used in this paper refers to the model selection method that jointly estimates the optimal lag length and cointegration rank by minimizing the information criteria over a set of

[2] At this regard, recent literature (Gonzalo and Pitarakis, 2006; Goetz and von Cramon-Taubadel, 2008) points out that the common use of term "threshold cointegration" in connection with TVECM could be misleading. In fact, in TVECM is the error correction term that is subject to threshold effect, while the cointegration itself is assumed to be constant and linear.

[3] Andrews (1993) argued that setting π_0 between 0.05 and 0.15 is a good choice.

Table 1
F-test values for presence of seasonality.

Series	Variable without seasonal adjustment	Deseasonalized variables
	$F_{(11,138)}$	$F_{(11,138)}$
GasOil	7.267 (0.000)	0.056 (0.999)
RapOil	2.656 (0.004)	0.041 (0.999)
SunOil	2.517 (0.006)	0.166 (0.998)
SoyOil	2.080 (0.026)	0.036 (0.999)

Significance level in parentheses.

Table 2
Test for unit root and stationary.

	ADF	PP	KPSS
No trend			
GasOil	−1.841	−1.908	0.871***
DGasOil	*−12.369****	*−12.430****	*0.081*
RapOil	−0.654	−0.381	1.112***
DRapOil	*−11.285****	*−13.886****	*0.199*
SunOil	−0.374	−0.412	0.724**
DSunOil	*−12.085****	*−12.086****	*0.349**
SoyOil	−0.283	0.042	1.191***
DSoyOil	*−13.523****	*−13.724****	*0.319*
With trend			
GasOil	−2.752	−2.942	0.123*
DGasOil	*−12.343****	*−12.407****	*0.074*
RapOil	−2.872	−2.770	0.088
DRapOil	*−13.908****	*−14.053****	*0.056*
SunOil	−1.667	−1.670	0.292***
DSunOil	*−12.290****	*−12.290****	*0.026*
SoyOil	−2.813	−2.734	0.278***
DSoyOil	*−13.629****	*−14.109****	*0.050*

(***), (**) and (*) indicate the reject of the null at 0.01, 0.05 and 0.1 significance levels.

combination with various lag orders and cointegration ranks. Testing the performance of the estimators through Monte Carlo simulation, Wang and Bessler (2005) provide evidence that when the sample size is larger than 100, as in our case, the one-step procedure performs as well as, and perhaps better than, the trace test for all model specifications including trend specification, correlation between the errors and moving average magnitude.

Table 3 reports the results for two-step procedure. Here, the data indicate the presence of one cointegrating vector in the RapOil–GasOil prices, thus these two series move jointly in the long term. The tests for SunOil–GasOil and SoyOil–GasOil indicate the absence of cointegrating vectors, suggesting that these two price pairs have a very unlikely long-term relationship.

As a matter of comparison, we report the results of the one-step procedure in Table 4. For RapOil–GasOil prices the Schwarz information criteria (SIC) are minimized at rank 0, indicating the absence of cointegrating relationships. On the other hand, the Hannan–Quin information criteria (HQ) are minimized at one lag and rank one indicating the existence of one cointegrating vector, consistent with the results of the two-

Table 3
Cointegration test.

Series	Hypothesized no. of CE(s)	Trace test	0.05 critical value	Prob.[a]
RapOil–GasOil	None	14.675	15.495	0.0662
	At most 1	0.071	3.841	0.7898
SunOil–GasOil	None	7.646	15.495	0.5041
	At most 1	0.435	3.841	0.5093
SoyOil–GasOil	None	5.999	15.495	0.6957
	At most 1	0.013	3.841	0.9092
Series	**Hypothesized no. of CE(s)**	**Max-Eigen test**	**0.05 critical value**	**Prob.[a]**
RapOil–GasOil	None	14.604	14.265	0.0442
	At most 1	0.071	3.841	0.7898
SunOil–GasOil	None	7.210	14.265	0.4646
	At most 1	0.435	3.841	0.5093
SoyOil–GasOil	None	5.986	14.265	0.6150
	At most 1	0.013	3.841	0.9092

Lags interval = 1 for the three series (selected by AIC, SC and HQ criteria).
Trend assumption: linear deterministic trend. Similar results were obtained with the other options.
[a] MacKinnon et al. (1999) p-values.

Table 4
Schwarz (SIC) and Hannan and Quinn (HQ) information criteria for one to four lags and from zero to rank two for the three pairs of market analysed.

Market pairs		SIC				HQ			
		One lag	Two lags	Three lags	Four lags	One lag	Two lags	Three lags	Four lags
RapOil–GasOil	Rank 0	−7.378[a]	−7.260	−7.198	−7.141	−7.450	−7.381	−7.367	−7.359
	Rank 1	−7.375	−7.257	−7.165	−7.109	−7.483[a]	−7.413	−7.370	−7.364
	Rank 2	−7.342	−7.224	−7.131	−7.075	−7.462	−7.392	−7.349	−7.342
SoyOil–GasOil	Rank 0	−7.381[a]	−7.237	−7.150	−7.085	−7.453[a]	−7.358	−7.319	−7.303
	Rank 1	−7.321	−7.180	−7.082	−7.026	−7.428	−7.336	−7.287	−7.280
	Rank 2	−7.287	−7.146	−7.048	−6.991	−7.407	−7.315	−7.266	−7.258
SunOil–GasOil	Rank 0	−6.923[a]	−6.784	−6.674	−6.615	−6.995[a]	−6.905	−6.843	−6.833
	Rank 1	−6.870	−6.729	−6.616	−6.565	−6.978	−6.885	−6.821	−6.820
	Rank 2	−6.840	−6.698	−6.586	−6.535	−6.959	−6.866	−6.803	−6.802

[a] Denotes minimum values.

Table 5
Test of weak exogeneity of GasOil and RapOil prices.

Prices	$\chi^2(1)$	Probability	Result
GasOil	1.940	0.164	Fail to reject
RapOil	6.245	0.012	Reject

step procedure. It has to be considered that for small to modest data set (i.e. 50–200 observations), SIC have be shown to be more susceptible to underestimate the number of cointegrating vectors or lags, while HQ are less prone to under-fit (Yu et al., 2007).

In our empirical analysis, it appears that RapOil–GasOil prices are cointegrated, as results from both the two-step procedure (Table 3) and the information criteria (Table 4) provide consistent results. For the other prices pairs (SunOil–GasOil and SoyOil–GasOil), both the one-step and the two-step procedures reject the presence of a cointegrating relationships.

In order to find which price is unresponsive to deviations from the long-run relations, we test for weak exogeneity (Engle et al., 1983) of the cointegrating price pair, RapOil–GasOil. The test is useful to identify the price series that evolves independently, corresponding to that which is weakly exogenous. Accordingly, the test identifies price leaderships, finding which price actually adjusts to maintain the long-run equilibrium (Asche et al., 1999). Following Johansen (1992, 1995), we tested for weak exogeneity on each series, testing every element of the adjustment matrix coefficient against zero. The likelihood ratio test is Chi-squared distributed with degrees of freedom equal to the number of cointegrating vectors. Results fail to reject weak exogeneity only for GasOil price (Table 5). Hence, as expected, GasOil shows evidence of prices leadership, and RapOil adjusts to long-run equilibrium as a consequence.

Since we are particularly interested in the analysis of the effect of biofuel policy implementation only on those markets that show a long-run relationship, following Balke and Fomby (1997), we then estimated a TVECM only for the cointegrated prices pair, RapOil–GasOil.[4]

The presence of a threshold was estimated via the application of the Hansen and Seo (2002) SupLM test (when β is estimated): this test supports the threshold hypothesis at a bootstrapped p-value of 0.0388. The estimated cointegrating coefficient was $\beta = -1.53$, showing strong responsiveness of the RapOil market to GasOil price movements. The estimated threshold value was $\gamma = -2.91$, and identified two regimes with statistically different ECM coefficients (the Wald test for equality for the ECM coefficient was significant at the 2% level). The first, or usual regime, occurs when $\text{RapOil}_t - 1.53 * \text{GasOil}_t \leq -2.91$, and included 82% of the observations, while the second, or unusual regime, included the remaining 18% of observations and corresponds to $\text{RapOil}_t - 1.53 * \text{GasOil}_t > -2.91$.

The estimated TVECMs are presented below (t-statistics are estimated using Eicker–White standard errors, reported in parentheses):

$$\Delta RapOil_t = \begin{Bmatrix} -0.177_{(-1.973)} - 0.059_{(-1.989)} w_{t-1} - 0.137_{(-2.035)} \Delta RapOil_{t-1} + 0.082_{(1.561)} \Delta GasOil_{t-1} + u_{1t}, \ w_{t-1} \leq -2.91, \\ -1.893_{(-2.458)} - 0.663_{(-2.447)} w_{t-1} - 0.266_{(-0.677)} \Delta RapOil_{t-1} - 0.043_{(-0.179)} \Delta GasOil_{t-1} + u_{2t}, \ w_{t-1} > -2.91, \end{Bmatrix}$$

$$\Delta GasOil_t = \begin{Bmatrix} 0.047_{(0.427)} + 0.016_{(0.453)} w_{t-1} - 0.076_{(-0.848)} \Delta RapOil_{t-1} + 0.025_{(0.319)} \Delta GasOil_{t-1} + u_{1t}, \ w_{t-1} \leq -2.91, \\ -0.883_{(-1.156)} - 0.322_{(-1.199)} w_{t-1} - 0.605_{(-1.548)} \Delta RapOil_{t-1} + 0.581_{(3.031)} \Delta GasOil_{t-1} + u_{2t}, \ w_{t-1} > -2.91. \end{Bmatrix}$$

In both the usual and the unusual regime, the GasOil adjustment parameters are not statistically significant, while the RapOil ECTs are significantly different from zero, so it can be hypothesized that the GasOil price drove the RapOil prices toward the equilibrium level. In the first regime, the magnitude of the RapOil ECT coefficient (−0.059) indicates slow adjustment to long-term equilibrium, whereas in the unusual regime the correction is 11 times faster (−0.663). Therefore, the convergence to long-term *equilibrium* was not uniform throughout the analysed period, i.e. it was faster when the deviation from equilibrium exceeded the critical threshold (see Fig. 1). Finally, the transitory effects expressed by the differenced terms highlight moderate autoregressive behaviour for RapOil prices in the usual regime, and more accentuate for the GasOil prices in the unusual regime.

[4] In the readers' interest, we would like to pinpoint that threshold cointegration analysis can be estimated for all price pairs (we thank an anonymous referee for suggesting this footnote). As we were aware of this essential point, initially we had also investigated the threshold-no cointegration hypothesis considered by Balke and Fomby (1997), applying the TVECM to the two non-cointegrating series (SoyOil–GasOil and SunOil–GasOil). However, we felt that searching for the presence of threshold type cointegration for the price pairs found to be not linearly cointegrated does not fit within the purpose of this study. The results of the test for the other two pairs of variables showed the existence of threshold cointegration only for SoyOil–GasOil (results are available upon request). Absence of linear cointegration and presence of the threshold cointegration point out that for SoyOil and GasOil the adjustment coefficients work just over the threshold value. This is not consistent with the theory, since fossil fuel is an input in the vegetable oil production chain and it is typically subject to high price fluctuations. Therefore, such 'local' behaviour could be simply interpreted as the adjustment cost in the vegetable oil production chain.

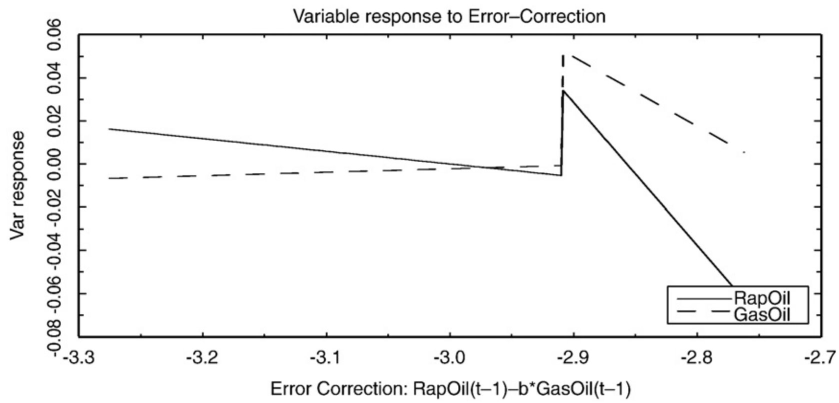

Fig. 1. Response of RapOil and GasOil to error correction, 2005:1–2007:11.

4. Conclusion

The growing concern about oil price rises and fluctuations, fuel security and environmental issues have led policy makers to remodel their energy portfolios. Within the wide range of renewable energy sources used to reshape the portfolio, a set of measure was developed in favour of biomass-based energy, and in particular for the biofuel segment. The aim of this paper was to investigate the presence of a long-term relationship between some vegetable oil and fossil diesel prices, as well as the existence of threshold effects.

During the period under consideration, in the EU, a political entity accounting for the production of 77% of world-wide biodiesel, the biofuel policies implemented have activated new price relations and dynamics. In order to analyse the issue, in this paper we present the results of a threshold approach (Balke and Fomby, 1997; Hansen and Seo, 2002) to support the non-linear nature of these price relations in the period 2005–2007, including price series for rapeseed oil, soybean oil, sunflower seed oil and fossil diesel. A threshold vector error correction model was used to consider short-run and long-run effects in two separate regimes.

Two main empirical results emerged from the empirical analysis. Firstly, cointegration analysis provided evidence to support that no long-run relationships exist among prices of sunflower seed oil and diesel and soybean oil and diesel, while rapeseed oil price is strongly linked to diesel price. More specifically, a test for weak exogeneity shows that the long-run behaviour of rapeseed oil is dependent on diesel prices and not vice versa. This is not surprising given the high quota (80%) of biodiesel produced by rapeseed oil in the EU, while other oils generally follow different production lines (e.g. food production). Secondly, threshold analysis, applied only to the price pair that is found to be cointegrated, showed the occurrence of asymmetric movements between these two prices, giving evidence of a threshold defining two different regimes. In extreme situations (18% of observations), the GasOil price drives the RapOil price to its long-run *equilibrium* in a stronger and faster way compared to the remaining periods. This non-linear relationship clearly indicates that rapeseed oil is more vulnerable to fossil fuel shocks compared to the past, as already pointed out by Yu et al. (2006) and Campiche et al. (2007).

The clear implication is that agricultural policy makers need to consider the new role of agricultural oil crops and their new emerging relationship with oil prices dynamics. This finding has important consequences, since the oil market has been shown to respond not only to global demand and supply conditions, but also to geo-political and institutional (i.e. OPEC) factors, as well as futures market dynamics (Sadorsky, 2004). Moreover, vegetable oil trader and biofuel producer, in order to reduce the risk related to price fluctuations are encouraged to manage other vegetable oils, with palm oil as the closest substitute (Liu, 2008). Importantly, the increasing demand for biofuel stimulates the conversion of land-use from undisturbed ecosystem to biofuel-related crops, causing a notable carbon debt that nullifies the environmental benefit generally associated with renewable energy sources (Fargione et al., 2008). In order to tackle the environmental issues that might arise as consequence of the increasing demand of other vegetable oils, the European Commission has recently proposed a regulatory framework for biofuel which requires the environmental sustainability of these crops (COM (2008)19 final).

Finally, the asymmetric property of the price relationship found in this paper could play a positive role in better understanding some economic properties of renewable energy policies focusing in the establishment and maintenance of an energy portfolio. In fact, as Awerbuch (2003) pointed out, the inclusion of a quota of renewable sources in the energy portfolio is a relevant strategy in managing risk. This is because even though fossil fuels could be cheaper than renewable sources, its high price fluctuations represent a risk for economic agents, hence a cost. On the contrary, the higher price stability of renewable sources of energy allows policy makers to manage less expensive energy portfolios. However, this theory, successfully tested to the case of renewable electricity generation (Awerbuch and Berger, 2003), is not likely to apply to biofuel from rapeseed. This outcome comes as a consequence of our findings, since the rapeseed oil market is subject to the same fluctuations of the market for diesel, with higher effects when these fluctuations exceed a certain threshold. Accordingly, in the EU the quota of energy produced by rapeseed biofuel does not seem to reduce risk in energy portfolio, and this should be considered in the analysis of the effectiveness of present and future energy policies.

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